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Abstract

The determinants of union membership in the UK are analysed using the BHPS (1991-2003). The paper employs three alternative methodologies to control for the problem of initial conditions. Trade union membership is found to be persistent even after controlling for the unobserved effect. There is evidence of a considerable correlation between the unobserved individual heterogeneity and the initial condition. Ignoring this overstates the degree of state dependence of union membership greatly. The extent of state dependence in union membership status is notably higher in the (1991-1996) period estimates and appears to be more pronounced in the case of male employees for the entire period under consideration. Concerning the observed heterogeneity the estimates suggest that an individual's propensity to unionise is determined by a mixture of industrial and personal characteristics that have a differential impact on male and female propensities.

Key Words: union membership, initial conditions, unobserved individual heterogeneity, state dependence, dynamic random effects probit models
Subject Classification: [JEL classification] C23, J51

1. INTRODUCTION

This paper analyses the determinants of union membership in the U.K during 1991-2003. We are interested in answering two questions:

i. What are the characteristics of the individuals who join trade unions?

ii. Is union membership a persistent phenomenon?

The legal framework governing collective bargaining arrangements in the United Kingdom was substantially altered during the 1980's. Stewart (1995) provides a concise account of the successive legislative changes that were targeted towards weakening the bargaining strength of UK trade unions.

The introduction of the 1980 and 1982 Employment Acts strengthened the case for claiming unfair dismissal on the basis of an employee's refusal to enter a closed shop. The 1982 Act dictated that all post-entry closed shops had to be sanctioned

[1]The views presented in this paper are the author's and do not reflect those of the BHPS data depositors, namely, the *Institute for Social and Economic Research* at the University of Essex, U.K.

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via a ballot of the workforce. The successive 1988 and 1990 Employment Acts prohibited all means to enforce a closed shop and thus, rendered post-entry and pre-entry closed shops, respectively, illegal. The overall effect was that the 1980s was a decade of a dramatic decline in aggregate trade union membership and union recognition in the UK (Stewart, 1995, pp.143-45).

Changes in the pattern of trade union membership and union recognition in the UK have been studied by a number of authors (e.g. Booth, 1986; Gregg and Naylor, 1993; Andrews and Naylor, 1994; Wright, 1995; Disney et al, 1996; Disney et al, 1999). Most of these studies employ cross-sectional data and are, therefore unable to control for unobservable individual specific effects something that might well result in biased estimates of the effect of observable attributes. Upon failing to let the effects of the explanatory variables be time dependent the effects of establishment characteristics are particularly prone to bias (see Arulampalam and Booth, 2000, p.290).

For instance, Booth (1986) using the 1975 National Training Survey, finds that establishment characteristics generally have a significant impact upon the probability of trade union membership while personal attributes, with a few exceptions[3], typically do not. Wright (1995)[4] suggests that gender does not differentially affect an individual's trade union membership decision although labour market experience is found to be positively associated with membership.

Arulampalam and Booth (2000) using data from the National Child Development Study was the first longitudinal study of the patterns of changes in individual membership across time for a single cohort of individuals (young men) in the UK. However, the nature of their data set does not allow them to discriminate between calendar time and life-cycle effects.

Given that closed shop practices have been effectively outlawed in Britain, individuals will voluntarily opt to enter the unionised sector only if the net expected return of union membership is positive and/or when management is supportive of union membership which is likely to occur under a non-adversarial style of industrial relations regime. On the other hand, provided that unions standardise wages for all workers within a sector an individual might as well not join, and thus not incur any costs such as an initiation fee, and instead free ride. In fact, the degree of free riding with regard to trade union membership is substantially larger in the UK as opposed to the US (see Arulampalam and Booth, 2000).

Costs and benefits of potential union membership are quite hard to quantify and the relevant data are not always available. Individuals might opt to become trade union members so as to benefit from incentive private excludable goods available only to trade union members. Arulampalam and Booth (2000) summarise these as protection against unfair dismissal, discrimination due to ethnic minority group membership, grievance procedures, pension plans advice and the implementation of well-defined dismissal arrangements in recessionary periods (p.291). Naylor and Raaum (1993) note that trade union membership can be "influenced by both social custom effects and by resources devoted by management to opposing unionisation" (p.591). Farber (1983) emphasises that "workers are heterogeneous in their preference for union representation to the extent that workers of different characteristics derive different amounts of pecuniary and non-pecuniary benefits from unionisation" (p.1420).

[3]Age and education for females and experience for males.

[4]Employs the 1986 Social Change and Economic Life Initiative (SCELI) dataset.

Unobserved individual heterogeneity plays an important role in modelling union membership status and failing to control for this provides biased estimates. However, it should be remembered that the existing data restrictions mean that we cannot attribute any specific effects entirely to unosberved heterogeneity. This study employs longitudinal data and by controlling for unobservable individual-specific effects it aims to assess the degree of state dependence of union membership and estimate the importance of observable and unobservable characteristics on union status determination.

The remainder of this paper is structured as follows. Section 2 outlines the data, sample selection and explanatory variables; Section 3 the estimation methodology; Section 4 presents and explores the estimated results; and Section 5 concludes.

2. DATA, SAMPLE SELECTION AND EXPLANATORY VARIABLES

The British Household Panel Survey (BHPS) originated in 1991 and follows the same representative sample of individuals over a number of time periods. The first wave of the Panel (1991-92) consisted of 5,500 households and 10,300 individuals drawn from 250 regions of Great Britain. In 1999 further samples of 1,500 households in the cases of both Scotland and Wales were added and in 2001 a sample of 2,000 households from Northern Ireland was included in the survey.

The data were split into two balanced panels (1991-1996 & 1997-2002) comprising of employees aged between sixteen and sixty five, excluding full-time students, that have participated in all of the respective six waves of the survey.[5] The ECHP[6] samples and the new Scotland, Wales and Northern Ireland samples were selected out.

The traditional practice in union literature is to use male manual full-time employees in order to estimate the determinants of union membership (see Swaffield, 2000, p.439). However, this study uses a full- time male employees' sample and a full sample of female employees (full-time and part-time inclusive of female employees on maternity leave).

Part-time male employees were excluded from the male regression sample used in this study in that the small gains in terms of sample size are more than outweighted by the costs of a potential increase in the heterogeneity of the male regression samples.

In addition, the full female sample can provide a comparison group against the male full-time employee sample that could potentially suffer from selectivity bias. According to Swaffield (2001) it might be possible, however, that the female sample is also prone to sample selection bias caused by the labour market participation decision (pp.439).

Further disaggregation towards a male manual full-time sample was not performed however as it would be fairly costly in terms of sample sizes. Descriptive statistics Tables for the respective selected samples are provided in Tables (II,III) in the Appendix and include of a set of personal and industrial characteristics.

[5] Proxy respondents allocated zero weights by the data depositors, by default, were selected out.

[6] European Community Household Panel survey.

2.1. Trade Union Membership

The dependent variable is an indicator function taking a value of one if an individual is a trade union member and a value of zero otherwise. An individual is taken to be a trade union member if he or she has responded positively to the question "Are you currently a member of: Trade Unions" in the Social and Interest Group Membership section of the BHPS. The respective variable in the BHPS data set is "Member of trade union". Unfortunately, this question was only asked every other year after the fifth wave (1995-96) of the survey since, the data depositors believe that there is not a lot of movement in and out of organisations and therefore, it was not felt necessary to ask this every year.

A further union membership variable which is available in the BHPS, "Member of workplace union", can be used as a proxy for union membership for waves six (1996-97), eight (1998-99), ten (2000-01) and twelve (2002-03). This variable is derived through the question that was asked conditionally following a positive response to the question "Is there a trade union, or a similar body such as a staff association, recognised by your management for negotiating pay or conditions for the people doing your sort of job in your workplace?".[7]

Following this response, a positive reply to "Are you a member of this trade union/association?" is recorded as membership of workplace union which includes "in-house" staff associations, but excludes employers' organisations. This introduces a degree of discontinuity in the BHPS data regarding trade union membership as this broader definition is also inclusive of staff associations, but, nevertheless when one wishes to undertake a longitudinal analysis spanning all of the first and last six waves of the BHPS this is the only alternative available.[8]

The "Union or staff association at workplace" variable allows us to construct a measure of union coverage. However, the fact that an individual's workplace is covered by a trade union that is recognised by management for bargaining purposes does not necessarily imply that this particular individual will actually be a member of the union.

According to Andrews et al (1998) while at the individual level members and non-members doing the same job within the same establishment earn the same wage, when comparing across establishments trade union membership is "a closer indicator of a differential than coverage"(p.453).

Thus, to conclude. The cross-sectional time series trade union membership variable constructed for all twelve waves of the survey suffers from a certain degree of discontinuity. Nevertheless, the two alternative questions available in the BHPS make it feasible to construct a measure of union membership. The individual responses from the two alternative union membership questions and the union coverage outcomes can be compared in order to detect any potential measurement error in union membership responses (see Swaffield, 2001). However, this is beyond the scope of this paper.

The descriptive statistics reveal that male union membership during (1997-2002) has fallen by 15.32 percent compared to the (1991-1996) levels. The respective female union membership rate has actually risen by 9.49 percent so that by the last cross section of the survey female unionisation rates converged to the level of male unionisation rates at a marginally higher percentage.

[7] The resulting variable is termed as "Union or staff association at workplace".

[8] Swaffied (2001) employs a similar approach.

2.2. Explanatory Variables

This section considers theoretical predictions, measurement issues and potential shortcomings with regard to the explanatory variables employed in the reported estimates. Table I displays the explanatory variables included in the trade union membership models. These are split into two categories[9], namely, individual and industrial characteristics.

Table I. Explanatory Variables

Personal Characteristics	Industrial Characteristics
Sex	Region
Employment Status	Size of Workforce
Potential Labour Market Experience	Industrial Classification (type of industry)
Marital Status	
Ethnic Background	
Highest Qualification Attained (education)	
Health Status	
Occupational Classification	

2.2.1. Personal Characteristics

It is generally accepted that male and female employees have different probabilities of unionisation. The theoretical prediction is that females will have a lower propensity to unionise due to their discontinuous labour market participation. The existing empirical evidence suggests that there has been a change in the gender composition of trade union members in Britain. Trade unions in the UK are currently more assiduous in their attempts to increase female membership rates. Budd and Mumford (2004) using the Workplace Employee Relations Survey (1998) show that trade unions seem to positively affect the provision of parental leave, special paid leave and job sharing options although a negative association was found with respect to flexible working hours options (p.220).

Commitments such as family care and maternity leave tend to disrupt female labour force participation and hence, the inclusion of the marital status and maternity leave variables in the female union membership models. Marital status is also included in the respective male equations since in the "exit-voice" scenario individuals with family obligations could seek greater job security through unionisation (see Booth, 1986, pp.46-47).

Employment status is also expected to affect an individuals' propensity to unionise. Theory predicts that part-time employees have a lower propensity towards unionisation owing to their less stable and discontinuous participation in the labour force (Booth, 1986, p.46). Hence, full-time[10] employment controls are included in the female union membership models along with maternity leave.

Labour market experience and age are expected to be positively associated with the probability of unionisation. Bain and Elsheikh (1979)[11] note that some trade unions have provisions for superannuation and sickness benefits thus, rendering unionism a more attractive idea for older employees (Booth, 1986, p.47).

[9] The grouping of explanatory variables adopted follows Booth (1986).

[10] "Full-time" is defined as employment requiring a minimum of 30 hours per week based on total hours (including both normal and overtime hours).

[11] In Booth (1986).

Disney et al (1999) employing data from the Family and Working Lives Survey[12] find significant differences in the age-union membership profile between cohorts over the past 20 years in the UK. Union membership was highly persistent and the probabilities of becoming a trade union member "have systematically declined across successive date-of-birth controls" (Disney et al, 1999, p.1).

To obtain a measure of potential labour market experience further education leaving age was subtracted from the individual respondents' age at the time of interview. Booth (1986) argues that using such a measure might actually over-estimate female labour market experience given the discontinuity effects of maternity leave and family care and instead proposes using age rather than experience.[13]

Ethnic minorities are also expected to have a higher propensity to unionise in order to protect themselves against unfair labour market discrimination. The "ethnic group membership" question of the survey was used to form the non-white ethnic background dummies included in the estimated models. These are "Black" (Caribbean, African and other), "Asian" (Indian, Pakistani, Chinese, Bangladeshi) and "Other Ethnic Minority Group".

Health status is a binary variable taking a value of one if the individual respondent has reported "Fair, Poor, Very Poor Self-Assessed Health Status". It is expected to have an adverse impact on the probability of active labour market participation and thus to be negatively associated with trade union membership.

During the ninth wave (1999-00) of the survey "General state of health" was provided in the place of "Health status over last 12 months" that was reported for all of the remaining eleven waves. Given the discrepancy in the wording of the self-assessed status question during the ninth wave of the survey, the corresponding categorisations were rectified accordingly to be brought in line with the remaining waves.[14]

Educational attainment is expected to have a differential effect on individuals' propensity towards unionisation. Abowd and Farber (1982) suggest that the standardisation of wage rates via the bargaining process implies that workers with a high degree of human capital invested in themselves would be less prone to be union members in that they would expect to receive reduced human capital premia within the unionised sector. Job queues for union jobs will therefore consist of predominantly low skilled workers while employers within the unionised sector will be aiming to employ those individuals with exactly the opposite characteristics in order to minimise costs.

The "Highest Academic Qualification" question of the survey was broken down to construct four distinct dichotomous qualification variables. These are, "First University Degree or Higher", "Vocational Qualifications: HND, HNC, Teaching", "A Level Qualifications" and "Secondary Education Qualifications: O Lever or CSE".

The Standard Occupational Classification 1990 (S0C)[15] is used to control for occupational attributes and nine major (one digit) groups were formed. The occupation of an employee can affect his propensity to unionise through the monopoly position granted by his occupation. Highly skilled employees can be valuable as-

[12] Cross-section of individuals between 16-69 years old in 1994/1995 with retrospective information on employment history.

[13] Estimations were performed with both measures and the respective results were fairly similar. In the light of this, all reported models are inclusive of potential labour market experience.

[14] See Hernández Quevedo, Jones and Rice (2004).

[15] Refer to Appendix 3.9 (BHPS), Standard Occupational Classification (SOC).

sets to employers and can thus have a strategic advantage over their low skilled counterparts when extracting concessions from firms (see Booth, 1986, p.48).

2.2.2. Industrial Characteristics

In the BHPS region, metropolitan area refers to the area of residence of an individual respondent and does not indicate the location area of the establishment since the BHPS is not a linked survey of individuals and establishments. Strictly speaking, this invalidates the classification of region as an industrial characteristic, although, the eventual grouping of the areas was sufficiently broad in order to preclude any significant commuting.

Larger establishments are more likely to enjoy product market power and this implies more quasi rents to be bargained over and a greater scope for unionisation. It is generally hypothesised that probability of unionisation rises with the size of the establishment.[16] Booth (1986) postulates that the size of the individual's workplace and the type of industry enter the union membership model through their impact on organisation costs. However, it might be the case that they are actually "proxying union bargaining power and social custom" (p.48). Disney et al (1996) treat union recognition as a function of plant size among other establishment characteristics under the hypothesis that establishment size may be acting as a proxy for union recognition.

The institutionalist view that large establishments act as if they were unionised and subsequently offer higher wages in order to avoid unionisation is of particular interest to this study in that this would reverse the positive association between union membership and the size of establishment. Brown and Medoff (1989) emphasize that if union avoidance efforts are a key part of the establishment size-wage relationship such an association must be less pronounced for workers who seem very unlikely to become union members. Nevertheless, the authors conclude that "however important union avoidance efforts may be, they are not an important part of the size-wage story" (p.1045). The BHPS provides a size of workforce variable and this was used to form four distinct size groups.

Stewart employing data from the Workplace Industrial Relations Surveys (1984, 1990) provides evidence that during the 1990s unions on average failed to achieve positive wage differentials and to gain recognition in establishments that were opened post-1984[17], whereas, for older (pre-1984) establishments the average 1990 union wage differential remained at 1984 levels (Stewart, 1995; pp.160-161). Disney et al (1996) using data from the three Workplace Industrial Relations Surveys (1980, 1984, 1990) also find a declining probability of trade union recognition in new establishments.

Unfortunately, unlike the Workplace Industrial Relations Survey (WIRS) the BHPS does not provide any information on the age of the establishment since is not a linked data set of workplaces and individuals[18] and therefore, this effect cannot be controlled for.

[16] Hirsch and Addison (1986) provide a comprehensive review of a number of studies establishing this.

[17] This effect was due to the introduction of the series of Employment Acts during the 1980s coupled with the shift in the equilibrium of control away from unions and in favour of management (see Stewart, 1995).

[18] On the other hand, Union membership in the Workplace Employee Relations (Panel) Survey (1990-98), new WIRS, can only be proxied by a construction of a union recognition variable.

Type of industry is controlled for by using the nine Industrial classification (one digit) divisions. These were constructed using the Standard Industrial Classification 1980 (SIC)[19] of the individuals' employer in his current employment post. Given that the coding frame was replaced by the Standard Industrial Classification 1992 from wave twelve onwards the latter was brought into line with the 1980 coding frame.

Traditionally, education and the public sector have been two of the most heavily unionised industries.[20] Further, the "Public Administration, Education, Other[21]" industrial classification is also associated with a higher proportion of individuals that are either educated to university level or holding vocational qualifications. The estimated models are therefore inclusive of interaction dummies between "Other Services SIC and University Qualification or Higher" and "Other Services SIC and Vocational Qualifications".[22]

3. ESTIMATION METHODOLOGY

The dynamic reduced form model depicting the decision of an individual to join either the unionised or non-unionised sector is outlined in equation (1) below. The benefits of employment within the unionised sector are captured by the latent variable y_{it}^*. The union membership status of an individual i in period t, is indicated by the dummy variable y_{it}.

The unknown parameters to be estimated are $(\gamma, \beta)'$ and x_{it} is a vector of exogenous explanatory variables (personal and industrial characteristics). The composite error term ν_{it} captures the unobserved individual heterogeneity underlying the union membership decision and is decomposed into an individual-specific component ε_i and an individual time-specific effect u_{it}:

$$\begin{aligned} y_{it}^* &= x_{it}'\beta + \gamma y_{i,t-1} + v_{it} \\ y_{it} &= I(y_{it}^* \rangle 0) \\ v_{it} &= \varepsilon_i + u_{it};\ t = 1, ..., T;\ i = 1, ..., N \end{aligned} \tag{1}$$

3.1. Potential Limitations

The combined effect of the 1988 and 1990 Employment Acts, that precede the introduction of the BHPS in 1991 was the effective outlaw of both post-entry and pre-entry closed shops (see Stewart 1995). This has the implication that individuals undertake their unionisation decision on the basis of wages, individual preferences and non-pecuniary benefits without coercion.

Potential seniority and non-pecuniary benefits can be sufficiently strong motives for individuals to remain within the unionised sector, irrespective of wage changes, and this introduces state dependence in the model (see Vella and Verbeek 1998). The obvious drawback, however, of adding the lagged union membership variable as

[19] See Appendix 3.7 (BHPS), Standard Industrial Classification 1980 (SIC).

[20] The male union membership percentages in the "Public Administration, Education, Other" industrial classification were {59.91, 57.97} during (1991-1996) and (1997- 2002) respectively. The corresponding figures for females are {45.85, 49.03}.

[21] "Other services" 1980 (SIC).

[22] The "Professional Occupations" classification is also characterised by a high proportion of highly skilled individuals. However, interacting this with the "University" and "Vocational Qualifications" controls has yielded neither statistically significant nor theoretically consistent results.

a regressor is that it gives rise to the problem of initial conditions which is discussed in greater detail in Section (3.2).

Of course, becoming a union member is also conditional upon the employer's hiring decision (see Abowd and Farber, 1982) and the primary deficiency of a reduced form union membership model is that it ignores the role of the employer in determining union status. Despite the fact that in the reduced form union membership model, employer attributes are captured through the industrial classification dummies these are not adequate in order to assign any specific effects purely to unobserved heterogeneity (see Vella and Verbeek, 1998; p.164).

The estimations presented here are inclusive of the size of establishment as an additional control concerning employer features, though as has already been mentioned in the previous section, the BHPS limits our ability to control for potential establishment age effects on union membership.[23]

Another potential shortcoming of the model concerns free riding on union bargained wages that apply to all employees within a covered establishment irrespective of their union membership status. Booth (1986) argues that wages can be viewed as a collective good by the individual therefore casting doubts on whether wages can actually be considered as a determinant of an individual's unionisation probability.

The collective good argument does not invalidate the theoretical prediction that union density is potentially positively associated with higher wages. At the margin an individual worker's union membership decision might not be affected by the level of wages set by collective bargaining. Primarily, this occurs because the individual does not consider that his decision will make an important difference to the aggregate union density in his sector. What is more, the "marginal" individual can always free-ride (see Booth, 1986, p.44 & pp.58-59).

Some readers might consider the exclusion of wages from the sectoral choice model of union membership as inappropriate. Wages can be included in the union membership models to detect whether individuals do free-ride and an insignificant coefficient can then validate the public goods theory of union wages given that earnings are not proxying any omitted variables. However, according to Booth (1986) this scenario is highly unlikely (p.44).

Vella and Verbeek (1998) suggest that the lagged value of union membership status affects an individuals' unionisation decision while it does not have any significant effect on the current wage. It is argued that union membership status may capture movement costs that are not specific to union employment. Workers are therefore assumed to change union membership status only if they change jobs. Furthermore, the long-term advantages of union employment, whilst generating persistence of union membership status, are not expected to have a significant impact on wages and therefore lagged union membership status is expected to have a minor effect on current wages (p.167).

The joint determination of union status and wages renders the coefficients of a single equation model inclusive of wages biased and inconsistent. Even if wages and union status are not determined simultaneously, which is rather improbable, wages might be acting as a proxy for omitted variables that are simultaneous such as job security and pension provisions and this would produce biased results. Since wages can be either a complement or a substitute of union negotiated non pecuniary benefits the direction of the bias cannot be determined a priori, thus rendering the interpretation of the resulting coefficients on wages problematic (Booth, 1986, p.

[23] Via the establishment "age effect" impact on union recognition (see Stewart, 1995).

43). A two-step methodology such as the one offered by Vella and Verbeek (1998, 1999b) provides a far more attractive alternative and this paper focuses on the first stage estimation of reduced form models of union membership determination.

3.2. Estimation Procedure and the Problem of Initial Conditions

Modelling any stochastic process with structural dependence among time-ordered outcomes requires initialising the process (Heckman, 1981a; pp.118). Initial conditions do become irrelevant asyptotically as $T \to \infty$ although, in the case of short panels as is the case in this study the problem cannot be overlooked since $T = 6$ and asyptotics instead rely on $N \to \infty$. The initial conditions problem occurs when the initial value of the dependent variable is correlated with unobserved individual heterogeneity. The presence of individual-specific effects ε_i clearly invalidates the assumption of exogeneity of union membership status in the first period of the survey.

The initiation of the stochastic process determining union membership has been in operation prior to initiation of the BHPS in 1991. This occurs since a large fraction of individuals in the samples used were labour market participants before 1991. Thus, the initial value of union membership cannot be taken to be exogenous. The conditional probability that an individual will become a union member in the future is a function of past experience.

The initial conditions problem cannot be readily ignored since the random effects maximum likelihood estimator in its standard form will be inconsistent (see Heckman, 1981a,c). Further, ignoring the correlation between individual-specific effects ε_i and the initial conditions will overstate the degree of state dependence.

State dependence and individual heterogeneity offer "diametrically opposite" explanations of the notion that those individuals who have experienced an event in the past are more likely to do so in the future (Hsiao, 2003, p. 216).

Considering otherwise identical individuals it is possible that, those who have experienced unionisation in the past will amend their preferences determining future propensity to unionise (e.g. via potential seniority and non-pecuniary benefits pertaining to union membership). This is an entirely behavioural effect.

Alternatively, it is possible individuals differ in specific unmeasured variables that affect their propensity to unionise while they are not influenced by experiencing unionisation per se. In the event whereby such variables are correlated over time, and are not appropriately controlled for, past experience may turn out to be a determinant of the individuals' future propensity to unionise since it acts as proxy for the temporally persistent unobservables. This is what Heckman (1981a, 1981c) terms as "spurious state dependence" as opposed to "true (structural) state dependence" occurring in the former scenario.

Union membership decision is modelled using the dynamic random effects probit specification given in equation (2). The random effects formulation was chosen since in the case of dynamic models with large N and small T fixed effects produce inconsistent estimates of the parameters as differencing out ε_i generates a linear regression model with lagged dependent regressors and serially correlated disturbances. Further, the random effects probit model is used instead of its logit counterpart since random effects yield correlations among the successive disturbances. For this purpose, the multivariate normal distribution is more flexible than the corresponding logistic distribution which requires that all correlations are equal to 0.5 (see Maddala, 1987):

$$y_{it}^{*} = x_{it}'\beta + \gamma y_{i,t-1} + \varepsilon_i + u_{it};\ u_{it} \sim N(0, \sigma_u^2) \tag{2}$$

The random effects model in its standard form assumes that ε_i is not correlated with x_{it}. The presence of the individual-specific time-invariant effect ε_i, however, renders the composite error term $v_{it} = \varepsilon_i + u_{it}$ temporally correlated even when the u_{it} are taken to be serially independent. Adopting the Mundlak (1978)- Chamberlain (1984) specification we can allow for a correlation between ε_i and the time means of the observed time varying characteristics taking the form of $\varepsilon_i = \overline{x}_i'a + \alpha_i$. Substituting this expression for ε_i in equation (2) we arrive at specification (3) where it is assumed that $\alpha_i \sim iidN(0, \sigma_\alpha^2)$ and is independent of (x_{it}, u_{it}) for all i and t:

$$y_{it}^{*} = x_{it}'\beta + \gamma y_{i,t-1} + \overline{x}_i'a + \alpha_i + u_{it};\ u_{it} \sim N(0, \sigma_u^2) \tag{3}$$

The individual-specific random effects framework suggests an equi-correlation, ρ, between any two successive disturbances for the same individual unit. Given that y_{it} is dichotomous, a normalisation is necessary and it is commonly assumed that $\sigma_u^2 = 1$. The resulting expression for ρ is given in equation (4):

$$\rho = cor(v_{it}, v_{is}) = \frac{\sigma_\alpha^2}{\sigma_\alpha^2 + 1}\ ;\ t, s = 2, ..., T; t \neq s \tag{4}$$

The initial conditions problem is tackled using three alternative estimation methodologies suggested by Heckman (1981c), Orme (2001) and Wooldridge (2005). Heckman's solution to the initial conditions problem approximates the reduced form marginal probability of the initial state by a probit function which has as its argument all of the available pre-sample information on the exogenous variables (Heckman 1981c, p. 188). Orme (2001) suggests a two step methodology which is an approximation if the correlation between the initial condition and individual random effects is weak. Wooldridge (2005) proposes an alternative approach which involves modelling the distribution of the unobserved effect conditional on the initial value and the observed history of strictly exogenous explanatory variables (p.40).

3.2.1. Heckman's Estimator

Stewart (2006a) provides a Stata® command, -redprob-, for Heckman's estimator of the dynamic random effects probit model. In the spirit of Heckman (1981c) we specify a linearised reduced form for the initial observation given by equation (5) where z_{i1} denotes a vector of strictly exogenous instruments such as pre-sample information[24] affecting an individual's propensity to unionise, the vector of means $\overline{x}_i$ and x_{i1}, η_i is correlated with $\alpha_i$but not with u_{it} for $t \succeq 2$:

$$y_{i1}^{*} = z_{i1}'\pi + \eta_i \tag{5}$$

In terms of orthogonal error components η_i can be expressed as:

[24] The "year current labour force status began" variable in the BHPS allows us to obtain information about the industrial/occupational classification for those individuals that have entered the labour force prior to 1991 and 1997 in the two panels. Further pre-sample information variables were constructed for the highest educational qualification of those individuals aged over 25 at the initiation of the respective panels.

$$\eta_i = \theta\alpha_i + u_{i1};\ \theta \succ 0 \tag{6}$$

By construction (α_i, u_{i1}) in equation (6) are independent of one another. Exogeneity of the initial conditions occurs when $\theta = 0$. Further, it is assumed that u_{i1} meets the same distributional assumptions as u_{it} for $t \succeq 2$. The linearised reduced form for sectoral choice in the initial time period is given by:

$$y_{i1}^* = z_{i1}'\pi + \theta\alpha_i + u_{i1};\ t = 1,\ i = 1, ..., N \tag{7}$$

The joint probability of $(y_{i1}, ..., y_{iT})$ for individual i, given α_i, suggested by Heckman's methodology is shown in (8) where Φ is the standard normal cumulative distribution function:

$$\Phi[(z_{i1}'\pi + \theta\alpha_i)(2y_{i1} - 1)].\prod_{t=2}^{T}\Phi[(x_{it}'\beta + \gamma y_{i,t-1} + \overline{x}_i'a + \alpha_i)(2y_{i1} - 1)] \tag{8}$$

Thus, expression (9) denotes the likelihood function to be maximised for a random sample of individuals. F stands for the distribution function of $\alpha^* = \alpha/\sigma_\alpha$ and given the normalisation adopted $\sigma_\alpha = \sqrt{\rho/1-\rho}$. Stewart (2006a) provides a program for this maximum likelihood estimator whereby, assuming that α is normally distributed, the integral over $\alpha^*$is evaluated using the Gaussian-Hermite quadrature:

$$\prod_i \int_{\alpha^*} \{\Phi[(z_{i1}'\pi + \theta\sigma_\alpha\alpha^*)(2y_{i1}-1)].\prod_{t=2}^{T}\Phi[(x_{it}'\beta + \gamma y_{i,t-1} + \overline{x}_i'a + \sigma_\alpha\alpha^*)(2y_{i1}-1)]\}dF(\alpha^*) \tag{9}$$

3.2.2. Orme's Pseudo Maximum Likelihood Estimator

Orme (2001) offers a "pseudo maximum likelihood estimator" *PMLE* procedure that rectifies the standard *MLE* for the effect of previous event history. The linear specification in terms of orthogonal disturbances in equation (6) allows for the possibility of non-zero correlation $r = corr(a_i, \eta_i)$. Orme's two-step methodology is an approximation when r is local to zero (i.e. when the correlation between the initial condition and the random effect is small).

Following Arulampalam et al (2000) we start with equation (3). Equation (6) is redefined as $\alpha_i = \delta\eta_i + w_i$ where by construction (η_i, w_i) are orthogonal of one another, $\delta = r\sigma_\alpha/\sigma_\eta$ and $var(w_i) = \sigma_\alpha^2(1-r^2)$. Substituting this new specification for the random effect into equation (3) we arrive at:

$$y_{it}^* = x_{it}'\beta + \gamma y_{i,t-1} + \overline{x}_i'a + \delta\eta_i + w_i + u_{it} \tag{10}$$

We begin with the linearised reduced form for the initial observation equation (5). Assuming that (α_i, η_i) are bivariate normal then $E(w_i \mid y_{i1}) = 0$ and $E(\eta_i \mid y_{i1}) = e_i$ where:

$$e_i = \frac{(2y_{i1} - 1)\varphi(z_{i1}'\pi)}{\Phi(\{2y_{i1} - 1\}z_{i1}'\pi)} \tag{11}$$

is the generalised probit residual[25]obtained from equation (5). Assuming that u_{it} is independent of x_{it} we take w_i to be the typical random effects probit error component given that we replace η_i by its conditional expectation.

[25]See Giourieroux et al (1987).

The assumption of bivariate normality of (α_i, η_i) renders the error component w_i in the second stage random effects probit model in equation (10) heteroskedastic since:

$$var(w_i \mid y_{i1}) = \sigma_\alpha^2(1 - r^2\zeta_i^2);\ \zeta_i = \frac{\varphi(z_{i1}'\pi)}{\sqrt{\Phi(z_{i1}'\pi)\Phi(-z_{i1}'\pi)}} \tag{12}$$

Effectively then Orme's two stage methodology involves estimating an "artificial" random effects probit model which is augmented by the generalised probit residual obtained from the first stage linearised reduced form for the initial period equation (5) under appropriate normality assumptions.

The rationalisation of Orme's approach is grounded on the assumption of r being local to zero so that $var(w_i \mid y_{i1}) \cong \sigma_\alpha^2$. Note that since union membership is a dichotomous variable the probit model for the initial period needs the normalisation that $\sigma_\eta = 1$. Given the expression for δ in equation (10) we obtain $r = \delta/\sigma_\alpha$. Rearranging equation (4) we arrive at $\sigma_\alpha^2 \cong \rho/(1-\rho)$ and hence, an estimate of r can be obtained by the expression provided in equation (13) where δ is the coefficient on the generalised probit residual from equation (10):

$$r = \delta\sqrt{(1-\rho)/\rho} \tag{13}$$

The "pseudo log-likelihood" to be maximised includes lagged union membership but treats y_{i1} as exogenous. Augmenting the set of regressors by e_i "provides a test of, and an approximate control for, the initial conditions problem". Locally equivalent alternative methodology guarantees that the usual t-test of the coefficient of the generalised residual gives an asyptotically valid test of $r = 0$ (Orme, 2001; pp.6).

3.2.3. Wooldridge's Conditional Maximum Likelihood Estimator

Wooldridge's (2005) solution to the initial conditions problem specifies a distribution of unobserved individual heterogeneity conditional on the initial condition instead of obtaining the joint distribution of all outcomes of the endogenous variables.

We begin by specifying the distribution of the unobserved effect as:

$$\varepsilon_i | y_{i1}, x_i \sim N(\alpha_0 + \alpha_1 y_{i1} + x_i'\alpha_2, \sigma_\alpha^2);\ x_i = \{x_{i1}, ..., x_{iT}\} \tag{14}$$

where the $(1xT)$ row vector x_i contains all non-redundant explanatory variables in all periods under consideration.

The presence of x_i in expression (14) implies that we are not able to identify the coefficients on time-constant explanatory variables in x_{it} although time-constant explanatory variables can be included in x_i and this is the main disadvantage of Wooldridge's approach. Including time-constant explanatory variables in x_{it} will only increase the explanatory power of the model as it is not possible to separately identify the partial effect of time-constant variables from their partial correlation with the unobserved effect (Wooldridge, 2005, p.44).

Of course, the implication of adopting the Mundlak (1978)- Chamberlain (1984) device allowing for a correlation between ε_i and the time means of the observed time varying characteristics, $\varepsilon_i = \overline{x}_i'a + \alpha_i$, with the Heckman (1981c) and Orme (2001) approaches is that the presence of the $\overline{x}_i$ means that we cannot separately identify the effects of time-constant variables there, either.

The density $D(y_{i1},...,y_{iT}|y_{i1}=y_1,x_i=x,\alpha_i=\alpha)$ is given by:

$$\prod_{t=1}^{T}\{\Phi(x_t'\beta+\gamma y_{t-1}+\alpha_0+\alpha_1 y_{i1}+x'\alpha_2+\alpha)^{y_t}.[1-\Phi(x_t'\beta+\gamma y_{t-1}+\alpha_0+\alpha_1 y_{i1}+x'\alpha_2+\alpha)]^{1-y_t}\} \tag{15}$$

To find the joint distribution of $(y_{i2},...,y_{iT}|\ y_{i1}=y_1,x_i=x)$ we need to integrate out α_i. Integrating (15) against the Normal (0, σ_α^2) gives the likelihood function in expression (16) which is identical to the structure of the standard random effects probit model with the only difference that the explanatory variables at time t are $\{z_{it}\equiv(1,x_{it},y_{i,t-1},\ y_{i1},\ x_i)\}$. It is assumed that data are observed for each cross-sectional unit in all time periods although given specific sample selection mechanisms (16) can be employed for the subset of observations forming a balanced panel:

$$\int_{\mathbb{R}}\prod_{t=1}^{T}\{\Phi(x_t'\beta+\gamma y_{t-1}+\alpha_0+\alpha_1 y_{i1}+x'\alpha_2+\alpha)^{y_t}. \tag{16}$$
$$[1-\Phi(x_t'\beta+\gamma y_{t-1}+\alpha_0+\alpha_1 y_{i1}+x'\alpha_2+\alpha)]^{1-y_t}\}(1/\sigma_\alpha)\phi(\alpha/\sigma_\alpha)d\alpha$$

Essentially then, what Wooldridge (2005) suggests is that by adding y_{i1} and x_i as additional explanatory variables in each time period under consideration and using a computationally easy standard random effects probit software we can estimate $(\beta,\gamma,\alpha_0,\alpha_1,\alpha_2,\sigma_\alpha^2)$.

4. ESTIMATES

This study focuses on the Heckman (1981c) and Wooldridge (2005) estimators. Regarding Orme's (2001) two-step methodology, the estimates of the correlation between the initial condition and the random effect, r, were all in excess of 0.8. In the light of this, we cannot ignore the inherent heteroskedasticity of the residual component, w_i, in the second stage random effects probit model (eq.10) as it produces inconsistent parameter estimates. The estimated results are therefore, not reported here.

The requirement that r is local to zero under Orme's estimator is quite a stringent condition. Arulampalam et al (2000) looking at unemployment persistence obtain significantly lower values of r. We would normally expect union membership to be far more persistent than unemployment, in that the former entails long term benefits whereas long term unemployment reduces your employment probability. Therefore, the high correlation between the initial value and the random effect found here is not surprising.

The remaining estimated models of union membership employing Heckman's (1981c) and Wooldridge's (2005) estimators are provided in Tables (IV-VII) in the Appendix. In all reported models the null under the Wald statistic[26] for multiple exclusion restrictions is rejected and hence, all inclusive covariates are jointly statistically significant. Further, the estimated parameters generally possess the theoretically predicted coefficients.

The reported random-effects probit estimates were arrived at employing the adaptive Gauss-Hermite quadrature method to compute the log likelihood and its

[26] LR for the pooled probit estimates.

derivatives since the large group sizes and the sizeable within group correlations, ρ, in all of the estimated models made the non-adaptive quadrature approximation inaccurate. The models were estimated by increasing the number of quadrature points to 30 which is equivalent to increasing the degree of the polynomial approximation (see Butler and Moffitt, 1982).

The quadrature checks undertaken revealed that the coefficient estimates were nearly invariant[27] to the quadrature point variation. The polynomial approximation with 30 quadtrature points is therefore sufficiently accurate and the estimated results can be interpreted with confidence.

The major advantage of Wooldridge's estimator is its computational simplicity which reduces estimation time substantially compared to Stewart's (2006a) command used to implement Heckman's estimator. Wooldridge's estimator for $t > 1$ was extended by a standard probit estimator for $t = 1$ to enable comparability with Heckman's estimator, although the latter still produced a superior log-likelihood for all models under consideration. The estimates obtained using both approaches are reported here as this reinforces the argument that the correlation between the initial condition and the unobserved heterogeneity results in an over-statement of the extent of state dependence in union membership, more effectively.

Stewart's -redpprob- command reports a pooled probit model for $t > 1$, the estimates of the reduced form model for the initial observation[28] and the dynamic random effects probit estimates[29] for $t > 1$. The pooled probit estimator treats the whole sample as a large cross section therefore assuming away all cross period correlation. This restrictive probit estimator provides an initial consistent estimate of the parameters although it is inefficient (see Maddala 1987) and the respective estimates are only provided here to demonstrate the overstatement of the degree of state dependence. The null hypothesis under the likelihood-ratio test for ρ was rejected in each of the estimated random effects probit models implying that $\rho \neq 0$ and therefore, the reported pooled probit estimates for $t > 1$ are rejected in favour of the random effects probit estimates.

Regarding both Heckman's and Wooldridge's estimators, the proportion of the total error variation attributed to unobserved individual heterogeneity, ρ, was significantly higher in the (1997-2002) dynamic random effects probit models for both male and female employees as opposed to the respective (1991-1996) estimates. Further, the unobserved heterogeneity in the female random effects estimates generally constitutes a notably greater proportion of the unexplained variance of the composite error term than it does in the corresponding male estimates.[30]

[27] Relative differences in the coefficients were well below the subjective rule of thumb of a less than one percent variation.

[28] The estimates of the reduced form model for the initial period are not reported here as they are not of direct interest.

[29] The reported random effects probit estimates are inclusive of either a set of industrial or occupational controls as the there is a certain degree of overlapping among the two classifications. Unless one of the two sets of controls was excluded, the maximum likelihood functions using the -redprob- command would not converge due to this collinearity. The same exclusions were made in the Wooldridge estimates for comparison purposes.

[30] Apart from the (1997-2002) Heckman estimator where ρ was marginally higher in the model for male employees.

4.1. The Persistence of Trade Union Membership

Turning to the exogeneity of the initial conditions, in the case of the Heckman estimator this requires that $\theta = 0$. It is evident that exogeneity[31] is strongly rejected in all models with the exception of the male (1991-1996) model where it is rejected at the not so stringent 10% level of significance.

With regard to the Wooldridge estimator, the coefficients on the initial value of union membership enter all of the estimated models with particularly strong and statistically significant effects that are much greater in magnitude than the coefficients on the lagged value of union status. This indicates that there is a considerable correlation between the unobserved individual heterogeneity and the initial condition. This correlation becomes even more pronounced in the (1997-2002) estimates for both genders compared to the respective (1991-1996) estimates and it is slightly more accentuated in the male models.

The coefficient on the lagged value of trade union membership enters all of the estimated models with a generally strongly statistically significant coefficient, an outcome that indicates that there is positive state dependence in union membership status even after controlling for the unobserved effect. The prevalent correlation between the initial condition and the unobserved individual heterogeneity renders the pooled probit estimates inconsistent and the consequent over-statement of the extent of state dependence is quite clear.

The pooled probit and random effects probit estimates employ different normalisations. The former uses $\sigma_\nu^2 = 1$ and therefore gives an estimate of γ/σ_ν whereas the latter employing $\sigma_u^2 = 1$ reports γ/σ_u. For comparison purposes the random effects estimates have to be multiplied by the factor $\sqrt{1-\hat{\rho}} = \sigma_u/\ \sigma_\nu$ so that they are converted into γ/σ_ν (see Arulampalam, 1999). The rescaled coefficients[32] on lagged union membership status are reduced even further compared to the "inflated" pooled probit estimates and in the case of both female (1997-2002) models they are driven to statistical insignificance.

To obtain the magnitudes of state dependence the average partial, marginal, effects $(\hat{p}_j - \hat{p}_0)$ and predicted probability ratios $(\hat{p}_j/\hat{p}_0)$ were estimated using the following counter-factual probabilities that take y_{t-1} to be fixed at 0 and 1 and are evaluated at $x_{it} = \overline{x}$ (see Wooldridge, 2006; Stewart, 2006b):

$$\hat{p}_j = \frac{1}{N}\sum_{i=1}^{N} \Phi\{(\overline{x}'\hat{\beta} + \gamma_j + \overline{x}_i'\hat{a})\sqrt{1-\hat{\rho}}\};\ \hat{p}_0 = \frac{1}{N}\sum_{i=1}^{N} \Phi\{(\overline{x}'\hat{\beta} + \overline{x}_i'\hat{a})\sqrt{1-\hat{\rho}}\} \quad (17)$$

The average partial effects and predicted probability ratios provided at the bottom of Tables (IV-VII) in the Appendix are an estimate of state dependence of union membership. The Heckman and Wooldridge estimators generally produce similar magnitudes while the respective pooled probit values are substantially "inflated".

The extent of state dependence in union membership status is markedly higher in the (1991-1996) period estimates as opposed to the (1997-2002) estimates with

[31] Testing for $\theta = 0$ must take into account that it lies on the boundary of the parameter space (see Stewart 2006a) .

[32] $\{1.249,\ 0.988;\ 0.256,\ 0.326\}$ for the (1991-1996; 1997-2002) male models for the Heckman and Wooldridge estimators respectively. Similarly, the corresponding values in the female models are $\{0.663, 0.625;\ 0.090, 0.151\}$.

regard to both genders. This outcome implies that for British employees during the period under analysis the probability of remaining a union member has declined. Further, the degree of state dependence in male union membership appears to be more pronounced than the respective female dependence in both time periods under consideration. This is not surprising as the male samples employed consist solely of full-time employees and moreover, male labour market participation is not as discontinuous as it is for female employees.

The predicted probability ratios in the case of the Heckman and Wooldridge (1991-1996) estimates suggest that a male worker with a given set of observable and unobservable attributes is 3.1 and 2.4 times, respectively, as likely to be a union member at period t if he had been so at $t-1$. The corresponding probability ratios from the (1997-2002) estimates are 1.3 and 1.6.

A female worker possessing a given set of observable and unobservable characteristics is approximately 2 and 1.7 times as likely to be a union member at period t if she had been so at $t-1$ according to the respective predicted probability ratios from the Heckman and Wooldridge (1991-1996) estimates. Similarly, the respective ratios from the (1997-2002) estimates suggest that a female worker is 1.1 and 1.2 times as likely to remain a union member.

4.2. Observed Individual Heterogeneity (Full-Time Male Employees)

The remaining explanatory variables were included in the two variants of the random effects probit model in order to control for some observed heterogeneity and were divided into personal and industrial characteristics.

Regarding the (1991-1996) estimates potential labour market experience enters the Heckman random effects probit model with a positive and statistically significant effect. The interaction dummies of "University Degree or Higher" and "Vocational Qualifications" with the "Public Administration, Education, Other" industrial classification enter both variants with sizeable positive and statistically significant coefficients. This is what was expected *a priori* as the public sector still remains one of the most heavily unionised industries in the U.K and it is associated with a relatively high proportion of highly skilled employees.

The occupational classification category that is excluded is "Other Occupations".[33] The negative statistically significant effects in all categories excluding "Craft and related" and "Personal and Protective Service" occupations indicate the widespread decline in male unionisation rates during (1991-1996). The "Clerical and Secretarial" and "Sales" occupations enter only the Heckman variant with statistically significant coefficients. The greater magnitude and statistical significance in both variants of the coefficients on "Professional Occupations", "Managers and Administrators" and "Associate Professional and Technical" suggests that individuals in white collar occupations display a greater degree of union aversion. This outcome is not surprising as we expect employees within these occupations to have a disincentive to become union members. This arises in that trade unions reduce human capital premia via the standardisation of wage rates (see Abowd and Farber, 1982). Further, white collar occupations consist of groups of highly skilled employees that can be seen as acting as a *de facto* union on its own since they

[33] The "Other Occupations" classification consists of: "other occupations in agriculture, forestry and fishing", "other occupations in mining and manufacture", "other occupations in construction", "other occupations in transport", "other occupations in communication", "other occupations in sales & service" and "other occupations not elsewhere classified".

cannot be rapidly and costlessly replaced (see Blachflower et al, 1990).

Turning to the industrial characteristics, the omitted area from the regional controls is the Midlands and this was done because its unionisation rate lies approximately half way through the respective rates in the remaining regions of the UK. Scotland enters only the Heckman variant with a negative and statistically significant effect and this means that Scottish male employees were less likely to join a trade union as opposed to their male counterparts in the Midlands during (1991-1996).

The size of workforce at the individual's workplace enters both of the random effects (1991-1996) model variants with positive and rather strongly statistically significant coefficients in the cases of Workforce >500 and Workforce 100-499. This effect seems to increase symmetrically with establishment size in terms of coefficient magnitude and statistical significance as was expected *a priori* given that the excluded workforce size is that of Workforce <25 employees.[34] This result invalidates the institutionalist view that larger establishments engage in union avoidance efforts by acting as if they were unionised and offering higher wages. Instead it supports the notion that larger establishments are more likely to enjoy market power and hence offer a greater scope for unionisation since there are more quasi rents to be bargained over.

In the (1997-2002) dynamic random effects estimates for full-time male employees the "Black" ethnic minority group enters both specifications with a sizeable positive coefficient although, in the Wooldridge estimates it is only statistically significant at the 10% level. This is in line with the expectation that individuals belonging to ethnic minority groups might have a higher propensity toward unionisation so as to safeguard themselves against unfair labour market discrimination. However, the "Other Ethnic Minority Group" enters the Heckman model with a negative and significant coefficient.

The interaction dummies of "University Degree or Higher" and "Vocational Qualifications" with the "Public Administration, Education, Other" industrial classification enter solely the Heckman variant this time with positive statistically significant coefficients.

The control for having at least a University degree enters both variants with a negative and significant effect, though this occurs only at the 10% significance level in the Wooldridge specification. This outcome is in line with Abowd and Farber's (1982) model that predicts that workers with a high degree of human capital invested in themselves would be less prone towards unionisation as they would expect to receive reduced human capital premia.

The "HND, HNC or Teaching" qualification category enters both variants with negative coefficients possibly reflecting the lower representation of male employees in the teaching profession.

Considering the industrial characteristics, among regional controls Wales appears with positive and statistically significant coefficients in both specifications. The North West enters only the Heckman variant with a similar positive effect. This implies that male workers residing in these regions have a higher tendency towards unionisation as opposed to those living in the Midlands which is the excluded regional classification.

All three establishment size controls enter both models with significant and posi-

[34] The 25-99 establishment size control enters with a reduced coefficient and a significant effect only for the Heckman specification.

tive effects although establishments with 100-499 employees appear with marginally higher coefficients than those with more than 500. This outcome offers further support against the institutionalist view that reverses the positive association between establishment size and unionisation.

The omitted industrial classification category is "Energy and Water Supplies" due to its particularly high unionisation rate. All industrial classification controls enter both specifications with negative and statistically significant effects except the "Public Administration, Education, Other" category.[35] This classification is a cluster of industrial sectors that have traditionally been some of the most heavily unionised sectors in the British economy. Apparently, it does not seem to share the fate of the British manufacturing industry, formerly the stronghold of trade unions, in terms of rapid deunionisation.

4.3. Observed Individual Heterogeneity (Female Employees)

The time average of marital status has been included in the estimated models to capture the correlation between the unobserved effect and the variable throughout time. It enters both random effects probit estimators for the (1991-1996) period with similar positive and statistically significant effects. The weight of empirical evidence suggests a changing pattern in the gender composition of trade union membership in Britain. The positive coefficient on mean marital status may reflect active union policies tailored to the needs of the female population in their efforts to attract more female members in an era that is characterised by rapid deunionisation among the male labour force.

The time average of full-time employment enters the Heckman specification with a statistically significant coefficient that possesses the theoretically predicted positive sign. This is consistent with the notion that the omitted part-time group is less prone towards unionisation due to its potentially unstable labour market presence.[36] The time average of maternity leave also enters the Heckman specification with a sizeable positive significant coefficient.

The interaction terms of "University Degree or Higher" and "Vocational Qualifications" with the "Public Administration, Education, Other" appear in both specifications with positive and statistically significant coefficients. The notably sizeable coefficients on the Vocational Qualifications interaction is not at all surprising given the high representation of women in both the teaching profession and public administration. Further, holding only A Level qualifications enters the Heckman variant with a positive and significant effect.

With respect to the occupational controls "Managers and Administrators", "Clerical and Secretarial" and "Sales" all enter the Heckman specification with significantly strong negative coefficients. "Managers and Administrators" and "Sales" occupations appear in the Wooldridge variant with a similar effects although the latter is only significant at the 10% level.

Regarding the industrial characteristics, the Heckman specification suggests that females residing in Scotland, Wales and the North East during (1991-1996) seem to be more likely to be union members than their counterparts in the Midlands. Living in London and the South East enters the Heckman model with a negative coefficient significant at the 10% level. In the Wooldridge variant the North East

[35] The "Transport& Communication" industry is solely significant at the 10% level in the Heckman model.

[36] Costs of union membership could exceed the potential benefits for this group.

appears with a significant positive effect and Scotland does so at the not so stringent 10% level. These might be reflecting differences in regional industrial structures.

Finally, with the exception of medium sized plants with 25-99 employees in the Wooldridge variant, all of the remaining establishment size controls enter both specifications with statistically significant and positive coefficients therefore, invalidating the negative association prediction of the institutionalist view.

Moving to the (1997-2002) random effects estimates for female labour market participants the time averages of marital status and full-time employment enter the Heckman model with positive significant coefficients. Further, Asian females and those belonging to the Other Ethnic minority group category enter the Heckman specification with significant and sizeable positive coefficients.

Holding only A Level qualifications enters both models with a significant effect and seems to be positively associated with female union membership during the period under consideration. This is also the case with the "University Degree or Higher" interaction with the "Public Administration, Education, Other" industrial classification which appears in both specifications with a fairly sizeable coefficient.

To complete the discussion on the personal characteristics set of variables, none of the occupational classification dummies enters either of the random effects models with a statistically significant effect. This outcome probably reflects the fact that female unionisation rates throughout the period concerned have been increasing albeit not markedly.

Considering in turn the industrial attributes and primarily the regional controls, according to both models female employees residing in London and the South East along with those in East Anglia appear less likely to be trade union members compared to those living in the Midlands during the (1997-2002) period. Note that both coefficients in the respective Wooldridge estimates are only statistically significant at the 10% level. On the other hand, Wales enters both specifications with significant positive coefficients.

The size of workforce effects remain qualitatively the same as in the (1991-1996) female union membership model, although the respective (1997-2002) coefficients are of a greater magnitude and the size of the estimated coefficients increases symmetrically with workforce size. Medium sized plants with 25-99 employees in the Wooldridge variant are now statistically significant at the 10% level. Once more, the estimated coefficients on establishment size provide evidence that it is associated with union membership in a positive manner.

5. SUMMARY AND CONCLUSIONS

This paper studies the determinants of trade union membership in the UK during 1991-2003. The use of longitudinal data allows us to ascertain whether individual union membership at any point of time is a persistent phenomenon or instead a random process. By employing three alternative methodologies to control for the problem of initial conditions the results suggest that trade union membership is quite persistent even after controlling for the unobserved effect.

There is evidence of a considerable correlation between the unobserved individual heterogeneity and the initial condition and failure to control for this overstates the degree of state dependence of union membership substantially. This correlation becomes even more pronounced in the (1997-2002) estimates for both genders compared to the respective (1991-1996) estimates and it is slightly more accentuated in the male models.

The extent of state dependence in union membership status is notably higher in the (1991-1996) period estimates as opposed to the (1997-2002) estimates: an outcome that implies that the probability of remaining a union member, during the entire period under analysis, in the U.K has declined. Further, the degree of state dependence in male union membership appears to be more pronounced than the respective female dependence in both time periods under consideration. This is not surprising since the male samples employed in this study consist of full-time employees only and moreover, male labour market participation is not as discontinuous as it is for female employees.

The analysis undertaken signifies that the unobserved individual heterogeneity plays a critical role in modelling union membership status and failing to control for this gives biased estimates. Arulampalam and Booth (2000) make a similar statement and note that this is of course nothing new. It is consistent with social custom union theories that propose a plethora of union membership determinants such as commitment to unions and solidarity which are generally not observed by the econometrician.[37] However, this is also consistent with missing data on employer characteristics. It should be restated that while employer attributes are captured through the industrial classification and establishment size controls these do not suffice so as to assign any specific effects purely to unobserved heterogeneity.[38]

The observed heterogeneity estimates suggest that an individual's propensity to unionise is determined by a mixture of industrial and personal characteristics that have a differential impact on male and female propensities. This is at odds with earlier studies, such us Booth (1986) and Wright (1995), generally employing cross-sectional data[39] and suggesting that it is mainly the industrial characteristics that typically have a significant impact on the propensity to unionise.

Establishment size enters all of the estimated models with the most prominent observed heterogeneity effect, other than lagged trade union status, and is positively associated with union membership. This is in line with the notion that larger establishments are more likely to enjoy market power and hence offer a greater scope for unionisation since there are more quasi rents to be bargained over.[40]

Concerning the (1991-1996) male union membership estimates it is evident that there is a widespread decline in male unionisation rates and it appears that employees in white collar occupations display a greater degree of union aversion. Further, highly skilled male employees within the public sector are more likely to be union members. This outcome is reinforced by the respective (1997-2002) estimates revealing that the public sector does not seem to share the fate of the British manufacturing industry in terms of rapid deunionisation.

Married females appear to be more likely to be a trade union member during (1991-1996) possibly reflecting the active union policies tailored to the needs of the female population in their efforts to attract more female members in an era of declining male union membership. Highly skilled females within the public sector have a higher propensity towards unionisation according to the observed heterogeneity estimates for both periods. This reflects the high representation of women in both the teaching profession and public administration which still remain two of the most heavily unionised industries. Finally, given the increasing female unioni-

[37] See Arulampalam and Booth, (2000), pp.290&308.

[38] See Vella and Verbeek (1998), Arulampalam and Booth (2000).

[39] Hence failing to control for unobservable individual specific effects.

[40] However establishment size could be acting as a proxy for union recognition (see Disney et al,1996).

sation rates throughout the period under analysis, the (1997-2002) female estimates show no evidence of prevalent deunionisation across any occupational classification.

Future work could entail looking at the individual responses from the two alternative membership questions and the union coverage outcomes in the BHPS in order to detect any potential measurement error and its impact on the estimates of union membership determinants. Further, it could involve evaluating the sensitivity of the Heckman and Wooldridge estimators to the normality assumption on the unobserved individual effects by employing a discrete mass point distribution to model the unobserved heterogeneity (see Stewart, 2006b). These are major research tasks.

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Appendix

Table II. Descriptive Statistics (1991-1996)

Gender	Male		Female	
Variable	**Mean**	**Std. Dev.**	**Mean**	**Std.Dev.**
Trade Union Membership	**0.4099**	**0.4919**	**0.3173**	**0.4655**
Log (1+Potential Experience)	3.223	0.634	3.166	0.715
Marital Status	0.692	0.462	0.683	0.465
Full-Time Employment	_	_	0.672	0.469
Maternity Leave	_	_	0.011	0.104
Black (Caribbean, African, Other)	0.001	0.035	0.005	0.068
Asian (Indian, Pakistani, Chinese, Other)	0.014	0.116	0.008	0.090
Other Ethnic Minority Group	0.009	0.093	0.005	0.068
Inner/ Outer London and R of South East	0.300	0.458	0.300	0.458
South West	0.097	0.296	0.070	0.255
Midlands	0.167	0.373	0.158	0.365
Scotland	0.078	0.269	0.097	0.296
Wales	0.053	0.224	0.047	0.211
North West	0.110	0.313	0.112	0.316
North East	0.150	0.358	0.175	0.380
East Anglia	0.044	0.205	0.041	0.198
Other Services SIC and University Qualification or Higher	0.061	0.240	0.079	0.270
Other Services SIC and Vocational Qualifications	0.034	0.182	0.052	0.222
University Degree or Higher	0.158	0.365	0.109	0.312
HND, HNC, Teaching	0.082	0.274	0.074	0.263
A Levels	0.239	0.426	0.155	0.362
O Levels or CSE	0.333	0.471	0.426	0.495
Fair, Poor, V Poor Self-Assessed Health	0.157	0.364	0.199	0.399
Workforce >500	0.213	0.410	0.156	0.363
Workforce 100-499	0.303	0.460	0.228	0.419
Workforce 25-99	0.271	0.445	0.275	0.447
Workforce <25	0.213	0.409	0.340	0.474
Industrial Classification Dummies				
Agriculture, Forestry & Fishing	0.016	0.125	0.006	0.080
Energy and Water Supplies	0.051	0.220	0.007	0.085
Extraction of Minerals & Manufacture of Metals	0.052	0.221	0.018	0.134
Metal Goods, Engineering & Vehicles Industries	0.156	0.363	0.050	0.219
Other Manufacturing Industries	0.122	0.328	0.067	0.250
Construction	0.044	0.204	0.006	0.076
Distribution, Hotels & Catering (Repairs)	0.116	0.320	0.177	0.382
Transport & Communication	0.091	0.288	0.030	0.171
Banking & Finance	0.134	0.341	0.153	0.360
Public Administration, Education, Other	0.219	0.414	0.484	0.500
Occupational Classification Dummies				
Professional Occupations	0.124	0.330	0.107	0.310
Managers & Administrators	0.180	0.384	0.097	0.296
Associate Professional & Technical	0.108	0.311	0.124	0.329
Clerical & Secretarial	0.099	0.298	0.333	0.471
Craft & related	0.182	0.386	0.026	0.161
Personal & Protective Service	0.069	0.253	0.117	0.322
Sales	0.033	0.179	0.076	0.266
Plant & Machine Operatives	0.152	0.359	0.037	0.188
Other Occupations	0.054	0.227	0.081	0.273
Number of Observations	**4818**		**5172**	

Source: BHPS (1991-2003), ISER, Essex, SN:4967, June 2004

Table III. Descriptive Statistics (1997-2002)

Gender	Male		Female	
Variable	Mean	Std. Dev.	Mean	Std.Dev.
Trade Union Membership	**0.3471**	**0.4761**	**0.3474**	**0.4762**
Log (1+Potential Experience)	3.558	0.456	3.544	0.467
Marital Status	0.665	0.472	0.642	0.479
Full-Time Employment	_	_	0.709	0.454
Maternity Leave	_	_	0.013	0.113
Black (Caribbean, African, Other)	0.002	0.047	0.006	0.079
Asian (Indian, Pakistani, Chinese, Other)	0.013	0.113	0.013	0.111
Other Ethnic Minority Group	0.007	0.080	0.005	0.072
Inner/ Outer London and R of South East	0.269	0.444	0.282	0.450
South West	0.094	0.291	0.095	0.293
Midlands	0.192	0.394	0.172	0.377
Scotland	0.078	0.268	0.089	0.285
Wales	0.056	0.230	0.050	0.219
North West	0.102	0.303	0.111	0.315
North East	0.170	0.376	0.160	0.366
East Anglia	0.040	0.195	0.041	0.198
Other Services SIC and University Qualification or Higher	0.052	0.222	0.101	0.301
Other Services SIC and Vocational Qualifications	0.024	0.152	0.052	0.221
University Degree or Higher	0.169	0.375	0.133	0.340
HND, HNC, Teaching	0.087	0.281	0.078	0.268
A Levels	0.267	0.442	0.225	0.418
O Levels or CSE	0.345	0.475	0.409	0.492
Fair, Poor, V Poor Self-Assessed Health	0.181	0.385	0.206	0.404
Workforce >500	0.206	0.405	0.173	0.379
Workforce 100-499	0.297	0.457	0.213	0.409
Workforce 25-99	0.248	0.432	0.271	0.445
Workforce <25	0.249	0.432	0.343	0.475
Industrial Classification Dummies				
Agriculture, Forestry & Fishing	0.006	0.078	0.005	0.067
Energy and Water Supplies	0.025	0.155	0.007	0.083
Extraction of Minerals & Manufacture of Metals	0.053	0.225	0.014	0.116
Metal Goods, Engineering & Vehicles Industries	0.138	0.345	0.030	0.171
Other Manufacturing Industries	0.151	0.359	0.062	0.241
Construction	0.055	0.227	0.006	0.079
Distribution, Hotels & Catering (Repairs)	0.122	0.327	0.185	0.388
Transport & Communication	0.105	0.307	0.036	0.187
Banking & Finance	0.143	0.350	0.154	0.361
Public Administration, Education, Other	0.202	0.401	0.502	0.500
Occupational Classification Dummies				
Professional Occupations	0.098	0.297	0.106	0.308
Managers & Administrators	0.196	0.397	0.120	0.325
Associate Professional & Technical	0.110	0.312	0.141	0.349
Clerical & Secretarial	0.081	0.273	0.302	0.459
Craft & related	0.202	0.401	0.021	0.143
Personal & Protective Service	0.059	0.235	0.131	0.338
Sales	0.035	0.183	0.090	0.287
Plant & Machine Operatives	0.158	0.365	0.030	0.170
Other Occupations	0.063	0.243	0.059	0.235
Number of Observations	**5538**		**5760**	

Source: BHPS (1991-2003), ISER, Essex, SN:4967, June 2004

Table IV: Dynamic Random Effects Probit Models of Union Membership (1991-1996), Males

	Pooled Probit		Heckman		Wooldridge	
Variable	**Coef.**	**z**	**Coef.**	**z**	**Coef.**	**z**
Lagged Trade Union Membership	2.562	43.040	1.731	10.380	1.331	10.330
Trade Union Membership (1)	_	_	_	_	2.164	8.460
Log(1+Potential Experience)	0.117	1.850	0.349	2.800	0.068	0.610
Marital Status	0.039	0.200	-0.047	-0.210	-0.046	-0.200
Mean(Marital Status)	0.060	0.290	0.289	1.110	0.238	0.890
Black(Caribbean, African, Other)	-0.096	-0.370	0.161	0.360	-0.516	-1.110
Asian(Indian, Pakistani, Chinese, Other)	0.431	0.630	0.169	0.190	-0.266	-0.230
Other Ethnic Minority Group	0.119	0.370	0.369	0.680	-0.040	-0.070
Inner/ Outer London and R of South East	-0.094	-1.060	-0.203	-1.350	-0.203	-1.280
South West	-0.146	-1.260	-0.351	-1.750	-0.183	-0.900
Scotland	-0.269	-2.030	-0.439	-1.970	-0.366	-1.580
Wales	0.171	1.170	0.362	1.490	0.407	1.560
North West	0.095	0.870	0.291	1.510	0.053	0.280
North East	-0.071	-0.700	-0.015	-0.090	-0.221	-1.220
East Anglia	-0.183	-1.160	-0.284	-1.070	-0.253	-0.910
Other Services SIC&Univ. Qualification/Higher	0.604	3.860	1.133	4.260	0.727	2.850
Other Services SIC&Vocational Qualifications	0.584	2.850	0.940	2.830	0.882	2.480
University Degree or Higher	-0.079	-0.510	-0.142	-0.580	-0.088	-0.340
HND, HNC, Teaching	0.083	0.500	0.237	0.860	0.146	0.500
A Levels	0.057	0.580	0.180	1.100	0.119	0.690
O Levels or CSE	0.086	0.990	0.178	1.230	0.170	1.120
Fair, Poor, V Poor Self-Assessed Health	0.143	1.290	0.153	1.200	0.156	1.150
Mean(Fair,Poor,V Poor Self-Assessed Health)	-0.264	-1.600	-0.306	-1.280	-0.281	-1.110
Workforce >500	0.441	4.690	0.673	5.220	0.545	3.930
Workforce 100-499	0.377	4.360	0.507	4.420	0.413	3.240
Workforce 25-99	0.215	2.410	0.252	2.190	0.208	1.630
Professional Occupations	-0.560	-3.360	-0.831	-3.370	-0.777	-2.900
Managers & Administrators	-0.648	-4.240	-1.028	-4.310	-0.817	-3.300
Associate Professional & Technical	-0.446	-2.740	-0.747	-3.060	-0.526	-2.020
Clerical & Secretarial	-0.307	-1.940	-0.553	-2.370	-0.387	-1.550
Craft & related	-0.260	-1.790	-0.343	-1.650	-0.232	-1.000
Personal & Protective Service	0.028	0.170	-0.104	-0.420	0.233	0.870
Sales	-0.607	-2.560	-0.894	-2.770	-0.541	-1.560
Plant & Machine Operatives	-0.317	-2.150	-0.480	-2.210	-0.499	-2.110
Time Dummy 1993	_	_	_	_	-0.002	-0.010
Time Dummy 1994	-0.035	-0.450	-0.047	-0.530	-0.037	-0.340
Time Dummy 1995	0.057	0.720	0.040	0.450	0.075	0.680
Time Dummy 1996	0.133	1.670	0.140	1.560	0.211	1.890
Constant	-1.782	-6.420	-2.373	-4.890	-2.198	-4.590
ρ	_	_	0.479	4.110	0.449	7.146
θ	_	_	3.140	1.820	_	_
Average Partial Effect	0.800		0.462		0.375	
Predicted Probability Ratio	9.007		3.138		2.449	
Log-Likelihood	-1142.13		-1524.70		-1066.65	

Source: BHPS (1991-2003), ISER, Essex, SN:4967, June 2004

Table V: Dynamic Random Effects Probit Models of Union Membership (1997-2002), Males

	Pooled Probit		Heckman		Wooldridge	
Variable	**Coef.**	**z**	**Coef.**	**z**	**Coef.**	**z**
Lagged Trade Union Membership	2.412	42.410	0.697	4.470	0.569	4.670
Trade Union Membership (1)	_	_	_	_	3.680	11.500
Log(1+Potential Experience)	0.017	0.240	0.033	0.100	0.044	0.250
Marital Status	0.038	0.200	0.263	0.950	0.263	0.970
Mean(Marital Status)	-0.053	-0.270	-0.036	-0.090	-0.245	-0.770
Black(Caribbean, African, Other)	0.436	1.680	1.289	1.970	1.104	1.770
Asian(Indian, Pakistani, Chinese, Other)	0.058	0.110	1.197	1.520	-0.934	-0.590
Other Ethnic Minority Group	-0.123	-0.370	-1.612	-2.470	-1.481	-1.390
Inner/ Outer London and R of South East	-0.130	-1.540	-0.221	-0.880	-0.326	-1.500
South West	0.077	0.710	0.384	0.830	0.368	1.320
Scotland	0.062	0.510	0.512	1.730	0.151	0.490
Wales	0.345	2.630	1.633	4.600	0.732	2.240
North West	0.205	1.950	1.636	3.250	0.295	1.090
North East	-0.009	-0.100	0.525	1.640	-0.041	-0.180
East Anglia	0.034	0.240	0.718	1.840	0.061	0.170
Other Services SIC&Univ. Qualification/Higher	0.403	2.350	0.798	2.130	0.319	0.840
Other Services SIC&Vocational Qualifications	0.223	1.030	1.645	3.120	0.472	0.960
University Degree or Higher	-0.448	-3.300	-1.175	-2.690	-0.617	-1.880
HND, HNC, Teaching	-0.381	-2.650	-1.644	-3.520	-0.817	-2.270
A Levels	-0.010	-0.100	-0.021	-0.050	0.168	0.690
O Levels or CSE	0.000	0.000	-0.069	-0.190	0.204	0.880
Fair, Poor, V Poor Self-Assessed Health	-0.102	-1.030	-0.135	-0.990	-0.154	-1.130
Mean(Fair,Poor,V Poor Self-Assessed Health)	0.009	0.060	-0.347	-1.020	0.123	0.390
Workforce >500	0.604	6.760	1.168	6.260	0.902	5.370
Workforce 100-499	0.615	7.370	1.169	6.060	0.974	6.170
Workforce 25-99	0.356	4.050	0.641	3.670	0.548	3.460
Agriculture, Forestry & Fishing	-1.187	-2.300	-2.613	-1.970	-2.580	-2.200
Extraction of Minerals & Manufacture of Metals	-0.796	-3.700	-1.481	-2.830	-1.303	-3.160
Metal Goods, Engineering & Vehicles Industries	-0.941	-4.730	-1.157	-3.260	-1.086	-2.850
Other Manufacturing Industries	-0.826	-4.210	-1.157	-3.310	-1.189	-3.210
Construction	-0.876	-3.970	-1.276	-3.290	-1.104	-2.670
Distribution, Hotels & Catering (Repairs)	-1.122	-5.340	-1.767	-4.730	-1.500	-3.720
Transport & Communication	-0.475	-2.370	-0.645	-1.660	-0.770	-1.970
Banking & Finance	-0.837	-4.230	-1.343	-3.670	-1.129	-2.850
Public Administration, Education, Other	-0.351	-1.790	-0.199	-0.530	-0.280	-0.720
Time Dummy 1999	_	_	_	_	-0.359	-3.030
Time Dummy 2000	0.031	0.410	0.002	0.020	-0.178	-1.510
Time Dummy 2001	-0.031	-0.400	-0.031	-0.290	-0.202	-1.700
Time Dummy 2002	0.112	1.410	0.230	1.950	0.056	0.450
Constant	-1.102	-3.170	-1.615	-1.120	-2.149	-2.600
ρ	_	_	0.865	26.760	0.672	16.724
θ	_	_	1.164	6.030	_	_
Average Partial Effect	0.767		0.095		0.089	
Predicted Probability Ratio	9.737		1.310		1.592	
Log-Likelihood	-1258.97		-1595.28		-1078.95	

Source: BHPS (1991-2003), ISER, Essex, SN:4967, June 2004

Table VI: Dynamic Random Effects Probit Models of Union Membership (1991-1996), Females

	Pooled Probit		Heckman		Wooldridge	
Variable	**Coef.**	**z**	**Coef.**	**z**	**Coef.**	**z**
Lagged Trade Union Membership	2.095	37.920	1.083	10.150	0.909	9.010
Trade Union Membership (1)	_	_	_	_	2.072	10.700
Log(1+Potential Experience)	0.082	1.670	0.132	1.250	0.095	0.980
Marital Status	-0.115	-0.750	-0.120	-0.650	-0.157	-0.830
Mean(Marital Status)	0.352	2.090	0.621	2.570	0.589	2.500
Full-Time Employment	0.037	0.300	0.109	0.720	0.138	0.900
Mean(Full-Time Employment)	0.261	1.800	0.755	3.280	0.297	1.390
Maternity Leave	0.026	0.100	-0.049	-0.170	-0.074	-0.250
Mean(Maternity Leave)	0.857	1.340	2.765	2.050	1.638	1.350
Black(Caribbean, African, Other)	0.317	1.030	0.682	0.970	0.554	0.870
Asian(Indian, Pakistani, Chinese, Other)	-0.010	-0.030	0.447	0.570	-0.496	-0.720
Other Ethnic Minority Group	-0.171	-0.420	-0.483	-0.510	-0.690	-0.750
Inner/ Outer London and R of South East	-0.148	-1.750	-0.337	-1.860	-0.117	-0.700
South West	-0.071	-0.570	-0.188	-0.680	-0.055	-0.220
Scotland	0.215	2.030	0.524	2.220	0.397	1.850
Wales	0.258	1.980	0.641	2.170	0.395	1.480
North West	0.012	0.120	0.090	0.400	-0.098	-0.470
North East	0.254	2.860	0.608	3.030	0.384	2.120
East Anglia	-0.079	-0.510	-0.347	-1.000	0.046	0.150
Other Services SIC&Univ. Qualification/Higher	0.809	3.910	1.369	3.870	0.840	2.450
Other Services SIC&Vocational Qualifications	1.207	4.040	2.147	4.290	1.641	3.370
University Degree or Higher	-0.407	-1.960	-0.315	-0.830	-0.298	-0.820
HND, HNC, Teaching	-0.621	-2.200	-0.773	-1.600	-0.671	-1.430
A Levels	0.141	1.420	0.434	2.050	0.282	1.450
O Levels or CSE	0.089	1.180	0.252	1.540	0.147	0.980
Fair, Poor, V Poor Self-Assessed Health	0.121	1.390	0.129	1.240	0.122	1.150
Mean(Fair,Poor,V Poor Self-Assessed Health)	-0.253	-1.890	-0.357	-1.420	-0.247	-1.070
Workforce >500	0.340	4.230	0.534	4.020	0.402	3.020
Workforce 100-499	0.247	3.310	0.335	2.730	0.291	2.360
Workforce 25-99	0.183	2.620	0.209	1.910	0.139	1.260
Professional Occupations	-0.119	-0.800	-0.293	-1.070	-0.435	-1.620
Managers & Administrators	-0.359	-2.600	-0.599	-2.420	-0.556	-2.300
Associate Professional & Technical	-0.021	-0.160	0.046	0.190	-0.052	-0.220
Clerical & Secretarial	-0.212	-1.920	-0.434	-2.010	-0.322	-1.560
Craft & related	-0.006	-0.040	0.112	0.320	0.044	0.130
Personal & Protective Service	-0.028	-0.230	-0.028	-0.130	0.064	0.290
Sales	-0.356	-2.490	-0.679	-2.570	-0.430	-1.680
Plant & Machine Operatives	-0.297	-1.740	-0.369	-1.160	-0.327	-1.060
Time Dummy 1993	_	_	_	_	0.166	1.660
Time Dummy 1994	-0.109	-1.500	-0.127	-1.460	-0.024	-0.240
Time Dummy 1995	0.085	1.180	0.113	1.300	0.216	2.110
Time Dummy 1996	0.257	3.590	0.401	4.470	0.532	5.060
Constant	-2.086	-9.330	-3.080	-6.330	-3.160	-7.040
ρ	_	_	0.625	13.050	0.527	11.595
θ	_	_	1.098	6.540	_	_
Average Partial Effect	0.704		0.247		0.244	
Predicted Probability Ratio	6.407		1.964		1.712	
Log-Likelihood	-1446.08		-1758.06		-1331.85	

Source: BHPS (1991-2003), ISER, Essex, SN:4967, June 2004

Table VII: Dynamic Random Effects Probit Models of Union Membership (1997-2002), Females

	Pooled Probit		Heckman		Wooldridge	
Variable	**Coef.**	**z**	**Coef.**	**z**	**Coef.**	**z**
Lagged Trade Union Membership	1.995	38.460	0.235	2.260	0.298	2.920
Trade Union Membership (1)	_	_	_	_	3.303	12.950
Log(1+Potential Experience)	-0.015	-0.260	-0.176	-0.730	0.005	0.030
Marital Status	-0.010	-0.060	0.193	0.860	0.212	0.950
Mean(Marital Status)	0.257	1.500	0.710	2.150	0.441	1.550
Full-Time Employment	-0.067	-0.610	-0.022	-0.150	0.012	0.090
Mean(Full-Time Employment)	0.288	2.230	1.074	3.400	0.400	1.590
Maternity Leave	-0.025	-0.110	-0.148	-0.480	-0.131	-0.420
Mean(Maternity Leave)	0.226	0.410	1.689	1.040	0.854	0.540
Black(Caribbean, African, Other)	0.059	0.280	0.233	0.470	-0.112	-0.160
Asian(Indian, Pakistani, Chinese, Other)	0.843	2.040	4.272	2.920	1.973	1.500
Other Ethnic Minority Group	0.536	1.690	1.899	2.520	1.496	1.560
Inner/ Outer London and R of South East	-0.211	-2.830	-0.989	-3.830	-0.422	-1.890
South West	-0.148	-1.500	-0.532	-1.660	-0.449	-1.540
Scotland	0.168	1.750	0.498	1.520	0.426	1.480
Wales	0.322	2.700	1.303	2.960	0.728	2.020
North West	0.015	0.170	-0.322	-0.830	0.038	0.140
North East	0.128	1.570	0.353	1.340	0.365	1.540
East Anglia	-0.280	-2.030	-1.090	-2.260	-0.751	-1.850
Other Services SIC&Univ. Qualification/Higher	1.110	6.000	3.006	5.960	2.324	4.980
Other Services SIC&Vocational Qualifications	0.472	2.500	1.005	1.440	0.476	1.030
University Degree or Higher	-0.565	-3.020	-1.059	-2.140	-0.976	-1.980
HND, HNC, Teaching	-0.107	-0.630	0.249	0.300	0.602	1.320
A Levels	0.127	1.460	0.682	2.430	0.579	2.240
O Levels or CSE	0.089	1.150	0.247	1.010	0.370	1.600
Fair, Poor, V Poor Self-Assessed Health	0.017	0.210	-0.042	-0.380	-0.126	-1.130
Mean(Fair,Poor,V Poor Self-Assessed Health)	-0.079	-0.650	-0.267	-0.640	0.039	0.130
Workforce >500	0.430	6.010	0.781	4.670	0.705	4.700
Workforce 100-499	0.251	3.730	0.649	4.380	0.555	4.020
Workforce 25-99	0.187	2.960	0.297	2.320	0.221	1.790
Professional Occupations	0.080	0.580	0.307	1.100	0.119	0.420
Managers & Administrators	-0.312	-2.490	-0.266	-1.030	-0.305	-1.160
Associate Professional & Technical	-0.001	-0.010	0.176	0.700	0.115	0.450
Clerical & Secretarial	-0.181	-1.670	-0.023	-0.090	-0.154	-0.630
Craft & related	-0.137	-0.700	-0.109	-0.270	-0.146	-0.350
Personal & Protective Service	-0.069	-0.590	0.178	0.680	0.154	0.590
Sales	-0.308	-2.390	-0.421	-1.490	-0.304	-1.090
Plant & Machine Operatives	-0.236	-1.390	0.007	0.020	-0.204	-0.570
Time Dummy 1999	_	_	_	_	-0.672	-6.300
Time Dummy 2000	0.360	5.560	0.550	5.980	0.230	2.220
Time Dummy 2001	-0.302	-4.450	-0.142	-1.530	-0.494	-4.580
Time Dummy 2002	0.387	5.940	0.701	7.310	0.380	3.600
Constant	-1.638	-6.350	-2.767	-2.620	-3.275	-4.270
ρ	_	_	0.853	42.570	0.744	27.086
θ	_	_	0.729	8.190	_	_
Average Partial Effect	0.672		0.036		0.047	
Predicted Probability Ratio	4.157		1.080		1.214	
Log-Likelihood	-1766.94		-1980.20		-1485.09	

Source: BHPS (1991-2003), ISER, Essex, SN:4967, June 2004